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Abstract: Based on a theoretical monetary exchange-rate model in continuous time this paper establishes a sequential estimation framework which is capable of indicating central bank intervention in the run-up to a currency union. Using daily pre-EMU exchange-rate data for the countries of the current euro zone, we find mixed evidence of active pre-EMU intervention policies (so-called *institutional frontloading strategies*). Our estimation framework is highly relevant to economic and political agents operating in financial markets of the upcoming EMU accession countries.

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Abstract: Based on a theoretical monetary exchange-rate model in continuous time this paper establishes a sequential estimation framework which is capable of indicating central bank intervention in the run-up to a currency union. Using daily pre-EMU exchange-rate data for the countries of the current euro zone, we find mixed evidence of active pre-EMU intervention policies (so-called *institutional frontloading strategies*). Our estimation framework is highly relevant to economic and political agents operating in financial markets of the upcoming EMU accession countries.

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1 Introduction

Following the Maastricht treaty the European Monetary Union (EMU) started on January 1, 1999 with a core group of the following 11 countries (the so-called EMU first-wave member states): Austria, Belgium, Finland, France, Germany, Ireland, Italy, Luxembourg, the Netherlands, Portugal and Spain. In the meantime, four other countries have entered the euro zone, namely Greece on January 1, 2001, Slovenia on January 1, 2007 and Cyprus and Malta both on January 1, 2008. As an important technical stipulation accompanying the adoption of the euro, each EMU member state had to fix its exchange rate at its central parity from the European Exchange Rate Mechanism (ERM) on its EMU entrance date.

A key question for all EMU countries in the run-up to their respective EMU accessions was how to implement an optimal transition from the pre-EMU system of floating exchange rates into the fixed-rate system of the euro zone.[1] In this context, many politicians and economists advocated the implementation of so-called *institutional frontloading strategies*. The basic idea behind this policy design was to coordinate exchange-rate and monetary policies of the future EMU countries prior to the eventual exchange-rate fixing such that the following two objectives could be achieved: (a) The prevention of inherent speculative attacks, and (b) the credible steering of the exchange rates towards their final conversion rates in the case of severe misalignments prior to the fixed-rate system (see, among others, De Grauwe, 1996; Obstfeld, 1998).

In this paper we try to quantify the degree of such interventionist pre-EMU policies for the majority of the current euro-zone countries.[2] In order to achieve this, we take up and extend an econometric framework developed in Trede and Wilfling (2007). Technically speaking, their methodology consists in estimating a mean-reversion parameter from a continuous-time monetary exchange-rate model on the basis of discretely sampled observations via a maximum likelihood approach. In this paper, we basically adopt their approach, but extend the analysis in three respects. First, we apply the estimation procedure to all of the above-mentioned EMU exchange-rate series, that is we process a much larger data set.[3] Second, we plug the estimation technology in a

[1] Before January 1999, all EMU first-wave countries were members of the ERM and as such their currencies were kept in exchange rate bands of more or less wide band width (in general $\pm 15\%$ around the central parity). After the start of EMU on January 1, 1999, the former ERM was replaced by the ERM II, a target-zone system in which all participating currencies have central parities against the euro with fluctuation band width of $\pm 15\%$.

[2] In particular, we include all first-wave member states plus Greece and Slovenia in our econometric analysis, but exclude the small economies of Cyprus and Malta.

[3] Trede and Wilfling (2007) focus on the derivation of the econometric technique and only analyze the Greek drachma.

sequential analysis in order to make statistical inference about the dynamic evolution of interventionist policies in the run-up to EMU. Finally, we validate our empirical results by applying an exchange market pressure (EMP) model.

The key result of our study is that interventionist policies are likely to have played an important role in some EMU countries, but not in others. Our overall results should be of interest to at least two groups of economic agents. First, to the policy makers of future EMU accession countries who are faced with the implemention of an adequate monetary policy that aims at entering the euro zone under financial market tranquility. Second, the results of our sequential analysis should prove useful to financial market participants for whom knowledge about an active monetary policy stance might provide profitable information.

The remainder of our paper is organized as follows: Section 2 reviews the continuous-time monetary exchange-rate model which forms the theoretical basis of our estimation framework. The data, the econometric estimation technique, the empricical results and a robustness check are presented in Section 3. Section 4 offers some concluding remarks.

2 Previous results and preliminaries

In this section we present a theoretical exchange-rate model in continuous time which is capable of tracing the exchange-rate evolution in the run-up to a currency union and which potentially accounts for an increasing interventionist policy stance during this period. Essentially, our model represents a special stochastic version of the well-known class of monetary exchange-rate models with flexible prices which became popular at the beginning of the 1990s in the modeling of exchange-rate dynamics under specific exchange-rate regimes.[4]

To set up the theoretical exchange-rate model, we consider a world with two open economies under perfect capital mobility in which the political authorities of both countries decide to create a currency union on the future date t_S. On the analogy of EMU, the authorities therefore announce at date t_A (the announcement date) to irreversibly fix the presently floating exchange rate from the starting date t_S onwards at the specific parity $\overline{x}$ (the conversion rate at which both economies enter the currency union).

In order to derive an explicit exchange-rate path under such a time-contingent switch in exchange-rate regime from floating to fixed rates, it is convenient to consider the

[4]The seminal paper of this strand of literature is Krugman (1991) who presents a model of exchange-rate behavior in a permanent target zone. For an early overview of this class of models see Bertola (1994) and the literature cited there.

well-known monetary exchange-rate model with flexible prices in continuous time. In this model with rational expectations, the logarithmic spot rate at time $t, x(t)$, equals the sum of an exogenously given macroeconomic fundamental, $k(t)$, plus a speculative component representing the agents' expectations about future changes in the currency value:

$$x(t) = k(t) + \alpha \cdot \frac{E[dx(t)|\phi(t)]}{dt}, \quad \alpha > 0, \tag{1}$$

where $E[\cdot|\cdot]$ denotes the expectation operator conditional on the information set $\phi(t)$ which contains all information available to market participants at time t.[5]

Within the original multi-equation monetary flex-price model the exchange-rate Eq. (1) represents a reduced form and the macroeconomic fundamental $k(t)$ measures the effects of different variables such as real output differentials (between the domestic and the foreign economy), money supply differentials and stochastic shocks to money demand on the exchange rate. The positive parameter α represents the semi-elasticity of real money demand with respect to a short-term interest rate (see for example Froot and Ostfeld 1992).

Besides constituting the reduced form of the monetary flex-price model, Eq. (1) may simply be viewed as a standard valuation equation which emphasizes the asset-price character of the exchange rate (see among others De Grauwe et al. 1999). Within this general framework, we can think of the fundamental $k(t)$ as a collection of all economic and political factors that agents consider to be important for the valuation of the current exchange rate. Here, typical factors affecting $k(t)$ include central bank interventions and virtually any other economic activity which changes the relative supply-to-demand conditions of the home currency. In this setting, α represents a parameter weighting the fundamental component against the speculative motives for currency valuation.

An important feature of the exchange-rate Eq. (1) is that it allows us to model alternative types of monetary policy regimes in the run-up to the currency union. Technically, we achieve this by letting the exogenous fundamental $k(t)$ evolve over time as a specific stochastic process that is capable of capturing the main characteristics of the policy regime under consideration. In this paper, we consider an *institutional frontloading regime* with continuously increasing interventionist activity towards the entrance into the currency union at date t_S. Such an exchange-rate regime can be modeled by letting the fundamental $k(t)$ follow a scaled Brownian bridge with stochastic

[5] In Eq. (1), $E[dx(t)|\phi(t)]/dt$ is an abbreviation of $\lim_{s\to 0}\{E[x(t+s)|\phi(t)] - x(t)\}/s$. Since x denotes the log nominal exchange rate, $E[dx(t)|\phi(t)]/dt$ represents the expected (instantaneous) rate of change in the nominal exchange rate.

differential representation

$$dk(t) = \frac{\eta \cdot [\overline{x} - k(t)]}{t_S - t} dt + \sigma \cdot dw(t) \qquad \text{for } t \in [t_A, t_S), \tag{2}$$

with $\sigma > 0$ denoting the so-called instantaneous standard deviation and $dw(t)$ the increment of a standard Wiener process. Assuming $\eta \geq 0$, we see that the drift term $\eta \cdot [\overline{x} - k(t)]/(t_S - t)$ represents the force that keeps pulling the fundamental k towards the long-run exchange-rate target value $\overline{x}$ in case of a current deviation of the fundamental process from the target $\overline{x}$. At this point it is informative to recall that in our EMU framework the target value $\overline{x}$ coincides with the ERM central parity and thus the final conversion rate.

The modeling of target- (or mean-)reversion in the fundamental process $k(t)$, for which specification (2) is an example, is a well-established concept in the literature on exchange-rate dynamics (Froot and Obstfeld, 1992). In principle, there are many conceivable economic sources that might induce target-reversion in the fundamental process (such as specific institutional characteristics of the foreign exchange market and certain perceptions of foreign exchange traders). However, the traditional literature on exchange-rate dynamics under specific monetary regimes (for example under a target zone) ascribes the major portion of fundamental target-reversion to an interventionist policy stance (Delgado and Dumas, 1992; Lindberg and Söderlind, 1992; Svensson, 1992) and it is exactly this view that we adopt here.[6]

Keeping this in mind and noting that the drift term $\eta \cdot [\overline{x} - k(t)]/(t_S - t)$ is *ceteris paribus* increasing in absolute value proportional to the inverse of the remaining time until the currency union is implemented, we arrive at the following interpretation of Eq. (2): as $t \to t_S$ the strength of target-reversion approaches infinity with probability 1 whenever $\eta > 0$ so that the Brownian Bridge process in Eq. (2) consistently models an *institutional frontloading regime* with continuously increasing interventionist activity towards the currency union starting at date t_S. To highlight the role of the parameter η in the drift term, we from now on refer to η as the *intervention intensity*.

In conjunction with the Brownian bridge specification (2), the general law of exchange-rate evolution in Eq. (1) represents a stochastic differential equation. This can be solved by stochastic integration techniques and the imposition of adequate economic constraints which appropriately reflect the anticipations of foreign exchange market

[6] It should be noted, however, that it may prove difficult to empirically detect the degree of mean reversion in structural exchange-rate models derived from the monetary flex-price equation (1). See for example Iannizzotto and Taylor (1999) and Taylor and Iannizzotto (2001) who treat this issue for exchange rates in a target zone.

participants regarding the entrance of both countries into the currency union on date t_S at the conversion rate $\overline{x}$. Ruling out arbitrage opportunities at the moment of transition into the currency union, i.e. imposing the condition $\lim_{t \to t_S} x(t) = \overline{x}$ with probability 1, Trede and Wilfling (2007) derive the (bubble-free) solution to Eq. (1) as

$$x(t) = \overline{x} + \frac{k(t) - \overline{x}}{\alpha(t_S - t)^\eta} \cdot \int_t^{t_S} \mathrm{e}^{(t-r)/\alpha} \cdot (t_S - r)^\eta dr \qquad \text{for } t \in [t_A, t_S). \tag{3}$$

In the next section, we will estimate this structural exchange-rate equation on the basis of discretely sampled data for the present EMU countries.

3 Econometric analysis

3.1 Data

Our original data consist of daily spot exchange rates of all EMU first-wave currencies plus exchange-rate time series of the Greek drachma and the Slovenian tolar. All EMU first-wave member currencies are analyzed vis-a-vis the German mark (DEM) while the Greek drachma and the Slovenian tolar are both analyzed vis-a-vis the euro.[7] All exchange rates were compiled from the historical tables of the OANDA-FXTrade-Website and are daily averages of interbank rates recorded seven days a week.[8] Figure 1 depicts the eleven EMU exchange-rate series.

[Insert Figure 1 and Figure 1 (continued) here]

As indicated in the exchange-rate Eq. (3), the sampling period of each currency has to be chosen as the time interval $[t_A, t_S)$. For all EMU first-wave currencies t_S coincides with January 1, 1999, while for the Greek drachma and the Slovenian tolar t_S coincides with January 1, 2001, and January 1, 2007, respectively.

In contrast to the determination of t_S, it turns out to be far more difficult to find the empirical counterpart to the announcement date t_A from the theoretical model. The reason is that our theoretical model in Section 2 presumes that the announcement at date t_A generates news to market participants insofar as at that date the future entrance into the currency union (and thus the future switch in exchange-rate regime from floating to fixed rates) is announced and that this announcement comes as a

[7]It should be noted that the two EMU first-wave members Luxembourg and Belgium already shared a common currency, the Belgian franc, prior to the adoption of the euro.

[8]See the OANDA-website at http://www.oanda.com/convert/fxhistory (download from July 25, 2007).

surprise to market participants. In reality, however, agents in the foreign exchange market typically are well-informed and behave rationally. Consequently, agents must be supposed to anticipate a prospective future EMU entrance long before any official announcement. As a result, they incorporate this anticipation into their currency valuation schemes.

A feasible approach to overcoming this inconsistency between the theoretical exchange-rate model and the real-world information structure on foreign exchange markets is to reinterpret the announcement date t_A from the theoretical model as the so-called *date-of-first-notice*, that is as the date at which real-world foreign exchange market participants perceive a potential future EMU entrance for the first time. Wilfling (2009) adopts this view and elaborates an econometric framework which makes it possible to detect the *dates-of-first-notice* for the majority of the EMU currencies considered here. Technically speaking, this framework makes use of the theoretically well-grounded result that exchange rates are subject to switching volatility-regimes during the transition from floating to fixed exchange rates. Such volatility regime-switches, which stem from changes in financial markets' assessments of a country's prospective EMU participation (i.e. of its future entrance into a fixed exchange-rate system) can be captured satisfactorily by so-called Markov-switching GARCH models. An important feature of this class of time-series models is that they offer inferential techniques which are capable of locating the volatility regime-switching dates thus helping us to detect the empirical counterparts of the dates t_A (i.e. the *dates-of-first-notice*).

Wilfling (2009) covers exactly the same currencies as we do here. Table 1 displays the empirical announcement dates as they have been identified by his volatility regime-switching approach. However, his analysis fails in identifying t_A for the Austrian shilling, the Spanish peseta and the Slovenian tolar. As a result, we exclude these currencies from our econometric analysis below.

[Insert Table 1 here]

3.2 Estimation technique

We now address the estimation of the structural exchange-rate Eq. (3). We are primarily interested in the country-specific estimates of the mean-reversion parameter η which may provide important information on the degree of interventionist exchange-rate policies in the run-up to EMU.

We estimate the model parameters α, η and σ by maximum-likelihood as suggested by Singer (1998). This method is capable of coping with the non-stationarity of the driving stochastic process $k(t)$ as defined in Eq. (2). Our observations are n timings (our time unit is one year) and the corresponding exchange rates:

$$(t_1, x(t_1)), ..., (t_n, x(t_n)).$$

Following Trede and Wilfling (2007), the loglikelihood function conditioned on the first observation is

$$l_n(\alpha, \eta, \sigma) = \sum_{i=2}^{n} \ln(p(t_{i-1}, x_{i-1}, t_i, x_i; \alpha, \eta, \sigma)), \tag{4}$$

with $p(t_{i-1}, x_{i-1}, t_i, x_i; \alpha, \eta, \sigma)$ denoting the transition density from x_{i-1} to x_i between the dates t_{i-1} and t_i (where, for ease of notation, we write $x_i \equiv x(t_i)$). Since at each date t_i the exchange rate x_i is a linear function of the Gaussian fundamental process $k(t_i)$, the transition densities in the loglikelihood function (4) are also Gaussian. Moreover, linearity implies that the mean and the variance of x_i conditional upon the information set $\phi(t_{i-1})$ can be derived from conditional means and variances of the fundamental process $k(t_i)$, which according to Eq. (2) are given by

$$E[k(t_i)|\phi(t_{i-1})] = \left(\frac{t_S - t_i}{t_S - t_{i-1}}\right)^{\eta} k(t_{i-1}) + \left(1 - \left(\frac{t_S - t_i}{t_S - t_{i-1}}\right)^{\eta}\right) \overline{x}, \tag{5}$$

$$Var[k(t_i)|\phi(t_{i-1})] = \frac{\sigma^2(t_S - t_{i-1})}{2\eta - 1} \left(\frac{t_S - t_i}{t_S - t_{i-1}} - \left(\frac{t_S - t_i}{t_S - t_{i-1}}\right)^{2\eta}\right). \tag{6}$$

Using Equation (3) and defining

$$I(t, t_S) \equiv \int_t^{t_S} \mathrm{e}^{(t-r)/\alpha} \cdot (t_S - r)^{\eta} dr,$$

we obtain the following moments for the exchange rate:

$$E(x_i|\phi(t_{i-1})) = x_{i-1} + (x_{i-1} - \overline{x}) \left(\frac{I(t_i, t_S)}{I(t_{i-1}, t_S)} - 1\right), \tag{7}$$

$$Var(x_i|\phi(t_{i-1})) = \frac{\sigma^2 I^2(t_i, t_S)}{\alpha^2(2\eta - 1)} \left((t_S - t_i)^{1-2\eta} - (t_S - t_{i-1})^{1-2\eta}\right). \tag{8}$$

The transition density is given by

$$p(t_{i-1}, x_{i-1}, t_i, x_i; \alpha, \eta, \sigma) = \frac{1}{\sqrt{Var(x_i|\phi(t_{i-1}))}} \varphi\left(\frac{x_i - E(x_i|\phi(t_{i-1}))}{Var(x_i|\phi(t_{i-1}))}\right), \tag{9}$$

with $\varphi(u)$ denoting the probability density function of a standard normal variate. From all these elements, we finally obtain the loglikelihood function as

$$l_n(\alpha,\eta,\sigma) = -\frac{1}{2}\sum_{i=2}^{n}\frac{(x_i - E(x_i|\phi(t_{i-1})))^2}{Var(x_i|\phi(t_{i-1}))} - \frac{1}{2}\sum_{i=1}^{n}\ln(Var(x_i|\phi(t_{i-1}))) - \frac{n}{2}\ln(2\pi). \tag{10}$$

In order to find reasonable starting values for the optimization of the loglikelihood function (10), we apply a three-dimensional grid with respect to the parameters α, η and σ. For statistical inference on the parameter estimates, we use the final information matrix from the optimization procedure as the estimated covariance matrix.

In what follows, we refer to the estimation procedure just described as *non-sequential*, since we use the entire sampling period to estimate the intensity parameter η. This estimate sheds light on the question of whether or not we can measure a statistically significant target-reverting force that might have steered a currency towards its final conversion rate. However, it is conceivable that the level of the intensity parameter η changes throughout the sampling period, possibly indicating a modified intensity of interventionist activity. To account for this, we additionally apply our estimation procedure to a sampling window which continuously expands over time and subsequently refer to this approach as *sequential*.

To be more specific, within the sequential framework we enlarge the estimation window on a daily basis until the start of the currency union at date t_S. In technical words, we sequentially estimate the exchange-rate path (3) for the sampling periods $[t_A, t_n)$ with

$$t_n = t_A + \gamma + \Delta, t_A + \gamma + 2\Delta, \ldots, t_S, \tag{11}$$

where $\Delta = 1/365$, which is equivalent to one day, and the constant γ guarantees a minimal number of observations in each estimation window so that our estimation procedure finds a meaningful optimum of the loglikelihood function. Preliminary calculations show that this minimal number of observations should be equal to 100 implying that γ should be set equal to 99/365 in Eq. (11).

Under the assumptions of our model the maximum-likelihood estimators are asymptotically normally distributed. As the finite-sample distribution might deviate from normality we implement a bootstrap hypothesis-testing approach following MacKinnon (2007). For this, we denote the usual t-statistic of the test $H_0 : \eta = 0$ versus $H_1 : \eta \neq 0$ by $\hat{\tau}$. The p-value is the probability that the test statistic $\hat{\tau}$ is larger (in absolute value) than the estimate actually calculated from the original sample. The finite-sample distribution of the test statistic under H_0 is approximated by a parametric bootstrap. We simulate a large number ($B = 1000$) of synthetic exchange-rate paths conditional on the point estimators $\hat{\alpha}_{\eta=0}$, $\hat{\sigma}_{\eta=0}$ and $\eta = 0$. The three parameters α, σ

and η are then re-estimated from each synthetic sample. The bootstrapped t-statistics of the null hypothesis are τ_j^*, $j = 1, \ldots B$. Following MacKinnon (2007), we perform a two-tailed test[9] with the bootstrapped p-value estimated by

$$\hat{p}^*(\hat{\tau}) = \frac{1}{B} \sum_{j=1}^{B} \mathbf{1}_{\{x \in \mathbb{R}: |x| > |\hat{\tau}|\}} \left(\tau_j^*\right),$$

where $\mathbf{1}_A(\cdot)$ denotes the indicator function of the subset $A \subset \mathbb{R}$. These bootstrapped p–values are derived for every exchange-rate series.

3.3 Empirical results

Table 2 displays the non-sequential estimation results for the eight exchange-rate time series described in Table 1. The estimation results suggest a split among the eight currencies into two groups. The first group consists of the Belgian franc, the Dutch guilder, the Finnish markka, the French franc and the Greek drachma each of which exhibit significant point estimates for all of the three parameters α, η und σ. The Irish punt, the Italian lira and the Portuguese escudo, forming the second group, mostly exhibit insignificant point estimates for the parameters η and α.

[Insert Table 2 here]

Comparing the η-estimates of the currencies from the first group, we find indication of a low degree of interventionist activities for the Belgian franc ($\hat{\eta} = 0.541$), indication of a medium degree for the Dutch guilder ($\hat{\eta} = 1.003$),[10] the French franc ($\hat{\eta} = 1.777$) and the Greek drachma ($\hat{\eta} = 1.238$),[11] and indication of a high level of interventionist activities for the Finnish markka ($\hat{\eta} = 2.883$). The η-estimates of the three currencies from the second group (the Irish punt, the Italian lira and the Portuguese escudo) are not indicative of a significant degree of interventionist activities.

Next, we extend our findings from the non-sequential analysis. For this, we take a closer look at the dynamic properties of the intervention intensity parameter η by sequentially estimating the exchange-rate path (3) on the basis of expanding sampling periods $[t_A, t_n)$ as described in Eq. (11). Figure 2 displays the respective series of the

[9] We assume $\hat{\tau}$ to be symmetrically distributed around zero.

[10] For the Dutch guilder we report the bootstrapped standard error. For all other currencies the conventional standard errors are confirmed by the bootstrap analysis. The results of our bootstrap hypothesis-testing approach are available upon request.

[11] Our point estimate $\hat{\eta}$ for the Greek drachma differs slightly from the estimate given in Trede and Wilfling (2007) since we use seven instead of five exchange-rate observations per week.

last hundred sequential η-estimates (i.e. $t_n = t_S - 100/365, t_S - 99/365, \ldots, t_S$) for the Dutch guilder, the Finnish markka, the French franc, the Greek drachma, the Italian lira and the Portuguese escudo.[12] In order to assess statistical significance, we have added (pointwise) 95% confidence bands around the sequential point estimates $\hat{\eta}$ to each panel in Figure 2.

[Insert Figure 2 here]

The key insight of Figure 2 is that the currency-specific dynamics of the η-estimates often reveal successive phases with changing significance of the η-parameter. The Dutch guilder, for example, exhibits a significant η-parameter in the non-sequential estimation based on the entire sample. Yet, in the sequential analysis we find significant η-values only when considering the last week of the expanding sample. Here, the η-estimates range between 0.72 and 1.19. A case revealing a somewhat reversed pattern to the Dutch guilder is the Italian lira. For the Italian lira we find a non-significant η-estimate in the non-sequential estimation, but significant estimates in the sequential analysis on an interval covering a few weeks between mid-November and mid-December 1998 with estimated intervention intensities ranging between 0.71 and 1.68.

The Finnish markka and the Greek drachma exhibit qualitatively similar patterns with respect to the dynamics of their η-estimates. In the sequential analyses, the vast majority of the η-estimates are significant and show a tendency to increase over time. A similar (but less pronounced) behavior is also revealed by the French η-estimates.

Finally, the sequential η-estimates of the Portuguese escudo appear to be in close line with the findings in the non-sequential analysis. The sequential η-estimates are not statistically different from zero throughout the entire expanding estimation window.

3.4 Robustness check via an EMP model

In order to assess the validity of our empirical results from the previous section, we apply a simplified version of an exchange-market-pressure (EMP) model due to Weymark (1995, 1997). The general idea behind measuring exchange market pressure is to model money market disequilibria that arise from the excess demand or supply of the domestic currency. Generally, the pressure on the exchange rate can be relieved either by allowing the exchange rate to revalue or by changes in the stock of foreign exchange

[12]For the Belgian franc and the Irish punt our sequential approach frequently failed in finding a reasonable maximum of the loglikelihood function. Thus, we excluded both currencies from the sequential analysis.

reserves (assuming non-sterilized foreign exchange market interventions). Adopting the setup of Weymark (1997) this may be formalized as

$$\mathrm{EMP}_t = \Delta x_t + \delta \Delta r_t, \tag{12}$$

where the exchange market pressure at date t (EMP_t) is expressed as the weighted sum essentially consisting of the percentage change in the observed exchange rate (Δx_t) plus the change in foreign exchange reserves expressed as a proportion of the monetary base (Δr_t). The elasticity $\delta = -\partial \Delta x_t / \partial \Delta r_t$ constitutes a 'model-specific conversion factor that allows the disparate units in which exchange rate and reserve changes are measured to be combined into a single, composite measure of external imbalance' (Weymark, 1997).

Provided that $\mathrm{EMP}_t \neq 0$, Eq. (12) can equivalently be written as

$$1 = \frac{\Delta x_t}{\mathrm{EMP}_t} + \frac{\delta \Delta r_t}{\mathrm{EMP}_t}. \tag{13}$$

This last equation relates the two alternative proportions of exchange market pressure relieved by (a) the observed exchange-rate change Δx_t, and (b) by the intervention activity Δr_t of the domestic authorities. The term $\delta \Delta r_t / \mathrm{EMP}_t$ on the right-hand side of Eq. (13) is particularly important since the change in the foreign exchange reserves is solely determined by the activities of the domestic authorities. Therefore, Weymark (1997) defines her intervention activity index by

$$\omega_t \equiv \frac{\delta \Delta r_t}{\mathrm{EMP}_t}, \tag{14}$$

which, by virtue of Eq. (12), can explicitly be calculated as

$$\omega_t = \frac{\delta \Delta r_t}{\Delta x_t + \delta \Delta r_t} = \frac{\Delta r_t}{(1/\delta)\Delta x_t + \Delta r_t}. \tag{15}$$

A problem arises when calculating the intervention index ω_t from real-world data since the elasticity δ is not directly observable, but has to be estimated from the data. In the literature, this is frequently achieved by specifying a structural model from which first δ can be characterized and estimated by observable variables so that ω_t can ultimately be computed as in Eq. (15). In this paper, however, we apply a model-independent approach suggested by Eichengreen et al. (1995) according to which the elasticity δ formally obtains as

$$\delta = -\sqrt{\frac{Var(\Delta x_t)}{Var(\Delta r_t)}}, \tag{16}$$

on the basis of which we first estimate δ and then calculate the intervention index ω_t.

The intervention index ω_t can be interpreted as follows: When $\omega_1 = 1$ the central bank intervenes to keep the exchange rate fixed. When $\omega_t = 0$ the central bank refrains from intervening and lets the exchange rate float freely. Values between 0 and 1 indicate that the central bank has intervened in order to reduce the pressure on the exchange rate. Negative values ($\omega_t < 0$) signal the central bank's attempt to magnify exchange rate changes. When $\omega > 1$ the intervention pushes the exchange rate into the opposite direction as compared to what would have occurred without intervention. It should be noted that ω_t is very sensitive to the relative magnitudes of $\delta, \Delta x_t$ and Δr_t so that ω_t can take on extremely large absolute values. For this reason, we follow Jeisman (2005) and censor the intervention index by 2 when $\omega_t > 2$ and by -1 when $\omega_t < -1$.

[Insert Figure 3 here]

Figure 3 shows the monthly intervention indices derived by the EMP approach for all 8 EMU currencies described in Table 1. In addition to our exchange-rate data used in the previous sections, we processed monthly data for the foreign exchange reserves and the monetary base which we compiled from the International Financial Statistics (IMF).

The EMP results appear to be in line with the overall picture of our exchange-rate estimation results from the previous section. Our non-sequential estimation results in Table 2 exhibit a significant parameter η (the intervention intensity in our monetary exchange-rate model) for the Belgian franc, the Dutch guilder, the Finnish markka, the French franc and the Greek drachma. Interestingly, the intervention index ω_t from the EMP approach reveals a compatible pattern for these currencies. In the second half of 1998 (2000 in the case of Greece) many ω-values exceed the value 1 indicating that central banks seemed to have intervened in order to reverse movements of the respective exchange rates. In the case of the Dutch guilder the negative ω-value in December 1998 suggests that the central bank magnified the convergence of the guilder towards the EMU conversion rate. This finding is in line with the sequential η-estimates for the Dutch guilder in Figure 2, which indicate a significant intervention intensity during the last week of 1998.

For the Irish punt, the Italian lira and the Portuguese escudo our non-sequential estimation results from Table 2 do not indicate any evidence of interventionist activities in the run-up to EMU. For the Irish punt and the Italian lira this finding is confirmed by the EMP approach. Both currencies' ω-values mostly range between 0 and 1 during 1998 suggesting that the respective central banks made no significant attempt to magnify or to reverse the exchange-rate movements. In the case of the Portuguese escudo, however, our estimation framework and the EMP approach lead to slightly incompatible results. In contrast to our η-estimates from both the sequential and the non-sequential analysis, the Portuguese ω-values from the EMP model in the second half of 1998 suggest that the central bank may have attempted to magnify the movements of the escudo.

4 Conclusions

This paper estimates structural pre-EMU exchange-rate equations for the majority of the present euro-zone countries. On the basis of a continuous-time model of exchange-rate dynamics we estimate an explicit target-reversion parameter which we interpret as an indicator of an interventionist policy stance taken up by the monetary and political authorities. In order to bolster up inferential results we conduct a sequential estimation analysis. Overall, our empirical findings are indicative of active pre-EMU policy stances in various countries (pointing to the implementation of *institutional frontloading strategies*).

However, a word of caution seems to be in order. It should be noted that the target-reversion parameter η, which we interpret as a potential indicator of interventionist activities, may simply express the (speed of) convergence of a currency towards its final conversion rate shortly before the entrance into the currency union. But the speed of this exchange-rate convergence might have also been caused by other factors outside the authorities' control, for example by institutional characteristics of the foreign exchange market or the (time-varying) anticipations of foreign exchange traders and their removal of arbitrage opportunities at the moment of transition into the system of fixed exchange rates.[13]

As disentangling the effects of all potential factors is generally impossible from an econometric point of view due to data limitations, two alternative routes to mitigate this problem are conceivable. First, one could try to gather evidence of institutional frontloading strategies from the regular releases of the financial institutions involved

[13]Cf. the remarks on the exchange-rate equation (1) in Section 2.

(for example from central-bank reports etc.). However, financial institutions are not obliged and generally do refuse to publish the extent of interventions in the foreign exchange market or the exertion of other exchange-rate affecting instruments. Thus, we cannot be sure that institutional releases contain all necessary information and therefore have refrained from analyzing such reports here.[14]

Second, one could perform a robustness check using an alternative methodology to elicit information about the intervention intensity. We compared our estimation results with a simple EMP model to confirm the possibility of sustained intervention in the foreign exchange market. Here, the key finding is that the EMP model produces results which are largely compatible with our estimates of the target-reversion parameter η from our sequential and non-sequential procedures. Therefore, we interpret high η-estimates as a signal potentially indicating an interventionist policy.

Our estimation technology and in particular its embedding into the sequential framework should be of interest to several economic actors of upcoming EMU accession countries. As suggested above, policy makers often tend to be restrictive in providing prompt information on interventionist activities. Yet, information on such activities is of high value for traders in all sorts of financial and derivative markets, for example for valuing exchange-rate and interest-rate sensitive claims. To these agents our sequential analysis may signal an active policy stance supporting them in their financial decisions.

Apart from that, other actors may simply be concerned with the probability of a specific country qualifying for future EMU admission. At the end of the 1990s this kind of uncertainty about entering the currency union punctually as scheduled by the Maastricht treaty led to the construction of so-called *EMU probability calculators* (see Bates, 1999, for an overview and in-depth comparison between alternative methodologies). Since our sequential estimation procedure provides an alternative approach to extracting information on pre-EMU monetary policies from financial market data, it can be regarded as a complement to the EMU-probability-calculator strand of literature.

Acknowledgements

We are grateful to Mark P. Taylor and the anonymous referee for their helpful comments which greatly improved the paper. We also thank Martin T. Bohl for his constructive support during various stages of the project.

[14] Trede and Wilfling (2007) analyze the Bank-of-Greece's Annual Report 2000 and find that the Greek central bank indeed implemented institutional frontloading strategies consistent with high η-estimates.

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Figures and Tables

Figure 1: Exchange rates

Figure 1: (continued)

Figure 2: Sequential estimates for η (bold lines) plus 95% confidence bands

Figure 3: Monthly intervention indices ω_t derived by the EMP model

Figure 3: (continued)

Table 1: Currency-specific sampling periods $[t_A, t_S)$

Table 2: Estimation results of the non-sequential analysis

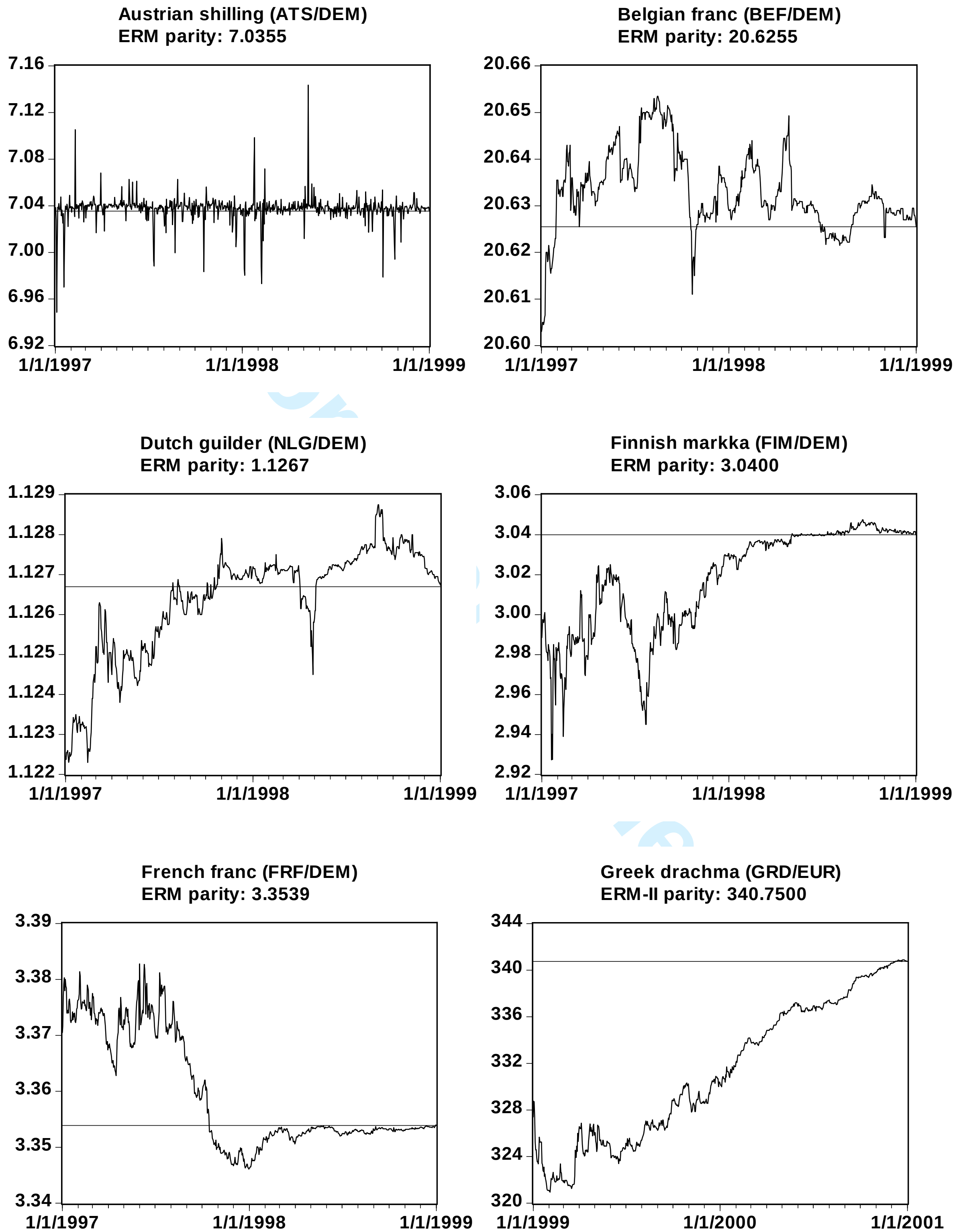


Figure 1: Exchange rates

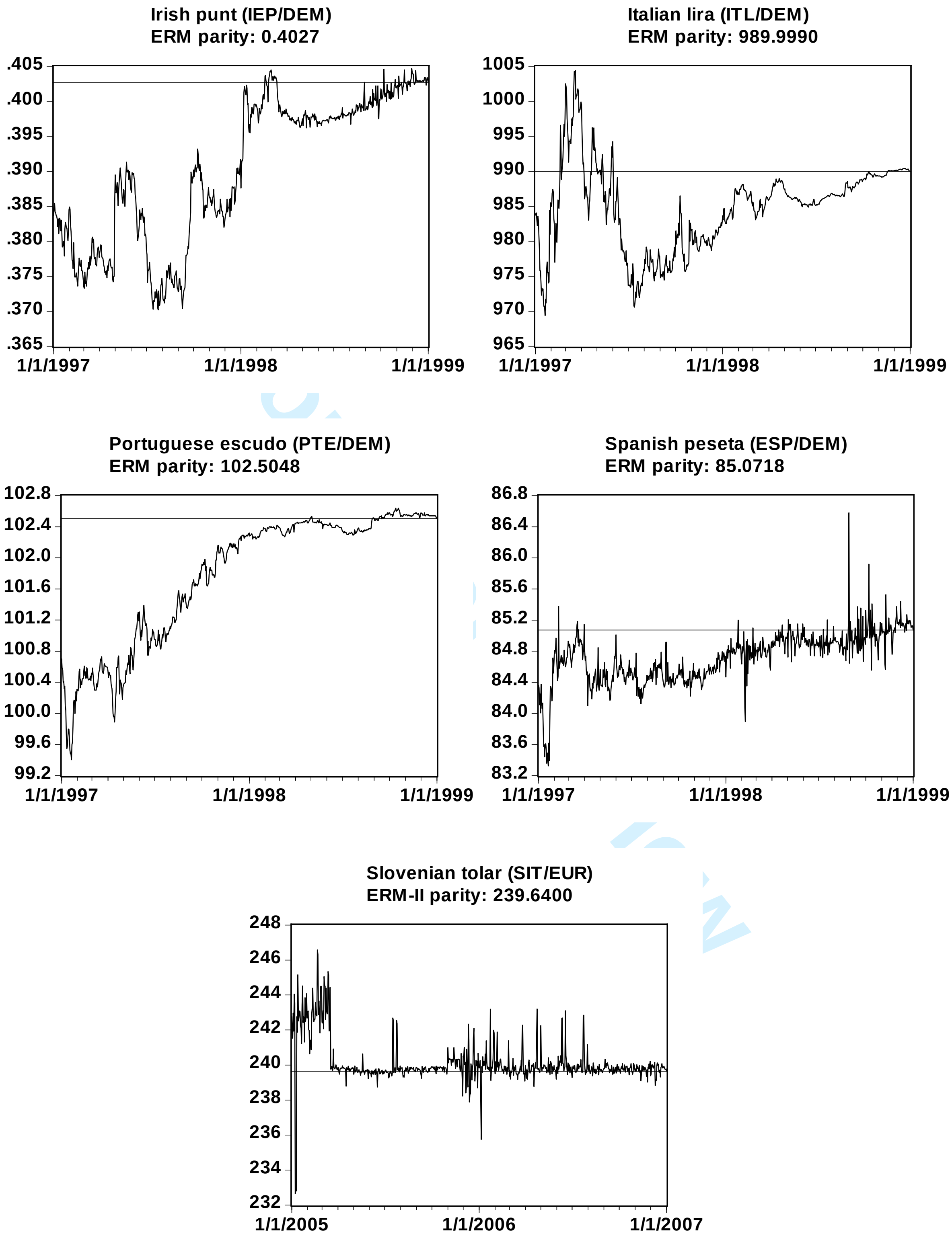


Figure 1: (continued)

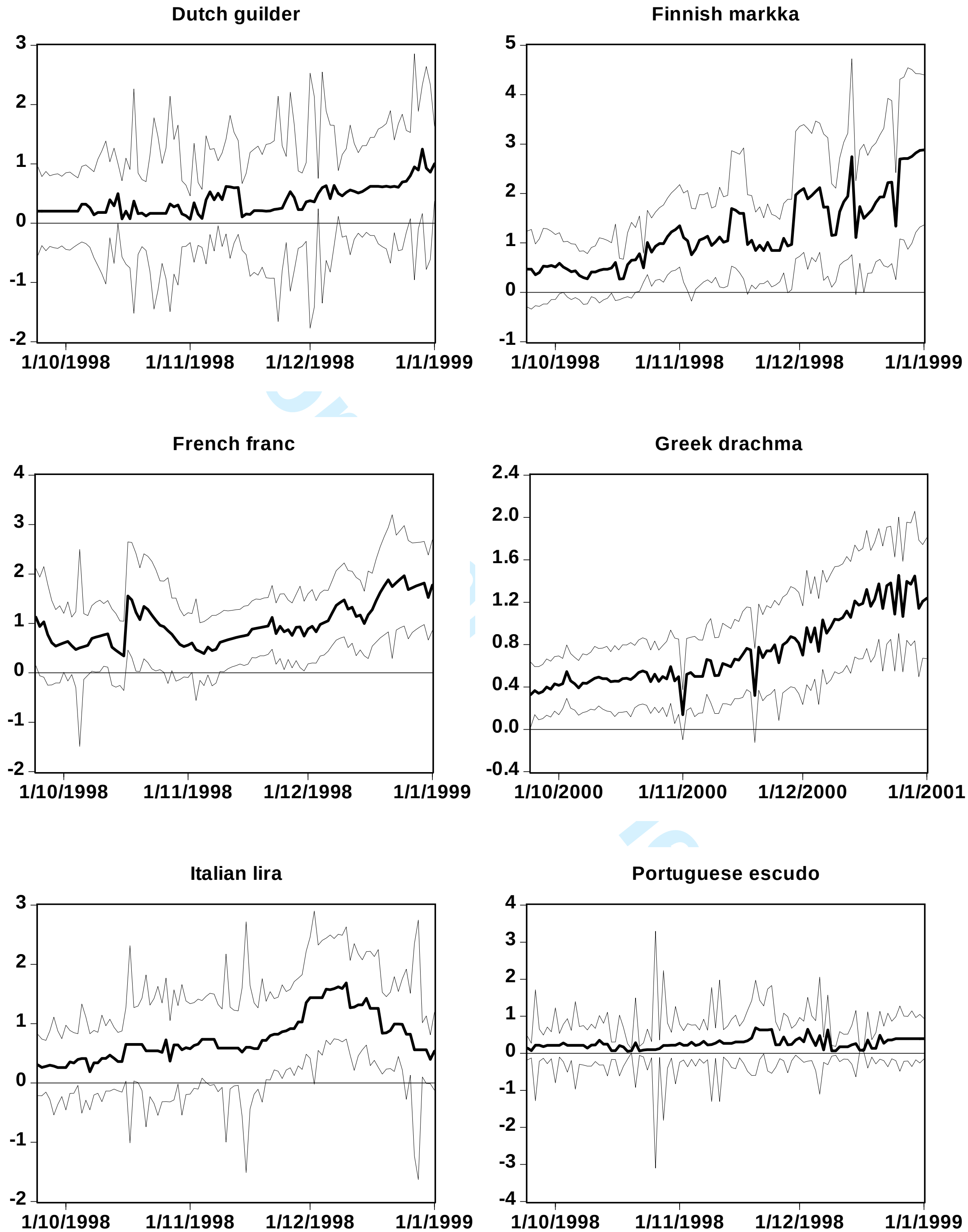
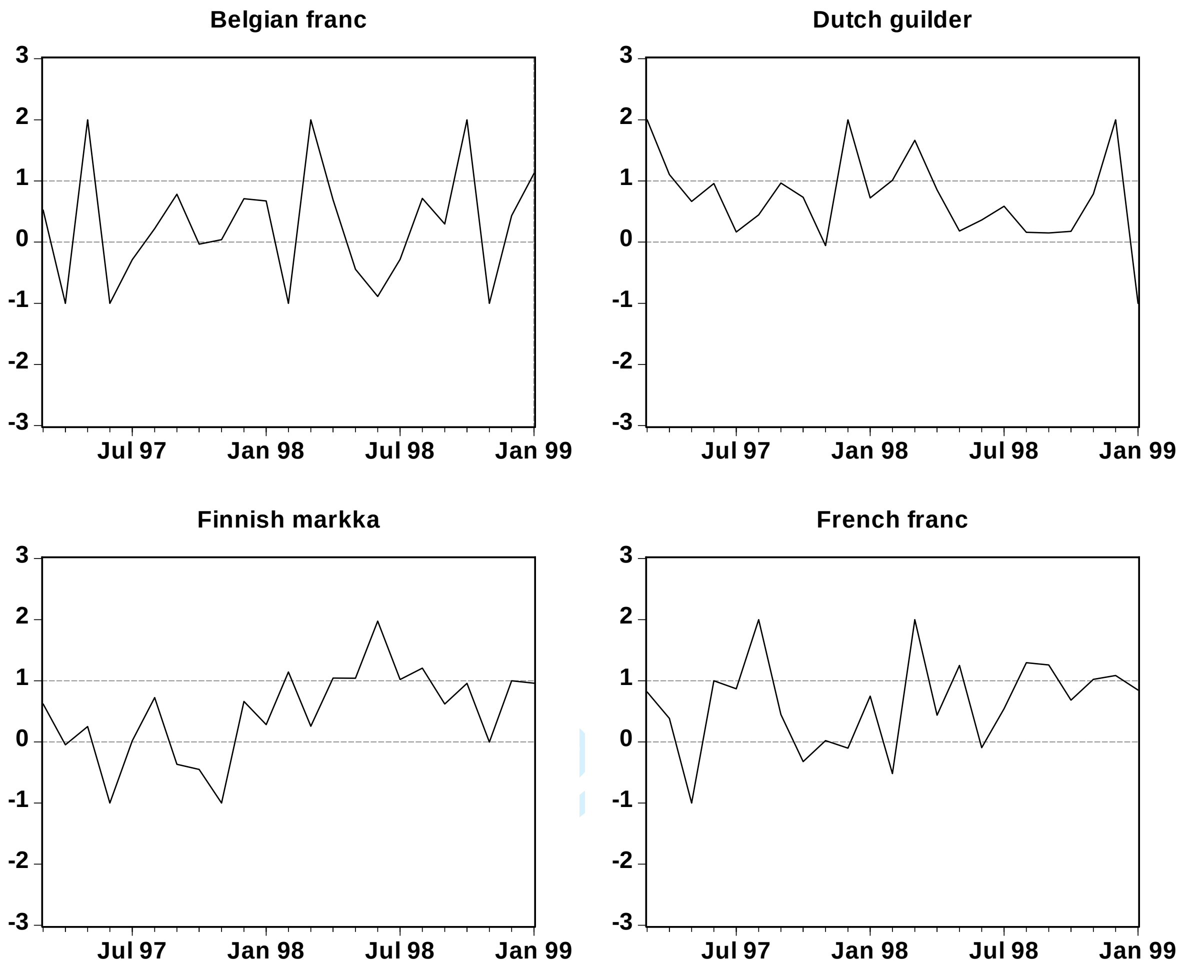


Figure 2: Sequential estimates for η (bold lines) plus 95% confidence bands



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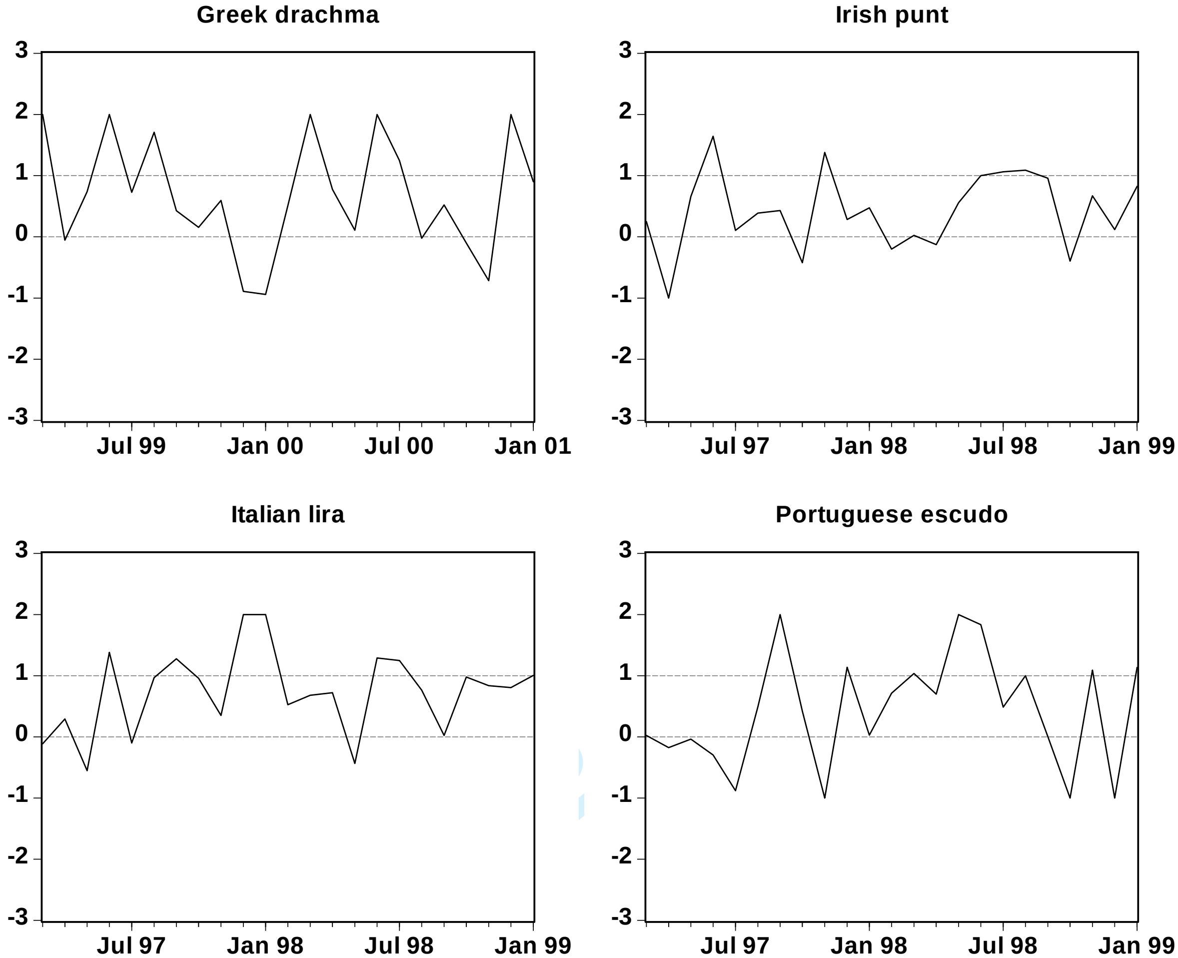

Figure 3: (continued)

Table 1: Currency-specific sampling periods $[t_A, t_S)$

Currency	Empirical announcement date (t_A)	Date of final conversion (t_S)	Number of observ. (n)
Belgian franc	May 5, 1998	January 1, 1999	243
Dutch guilder	November 5, 1997	January 1, 1999	423
Finnish markka	November 12, 1997	January 1, 1999	416
French franc	October 21, 1997	January 1, 1999	438
Greek drachma	March 1, 2000	January 1, 2001	219
Irish punt	May 20, 1998	January 1, 1999	227
Italian lira	April 4, 1998	January 1, 1999	275
Portuguese escudo	December 17, 1997	January 1, 1999	381

Notes : The empirical announcement dates are compiled from the preliminary analysis in Wilfling (2009).

Table 2: Estimation results of the non-sequential analysis

	α	η	σ
Belgian franc	0.003	0.541	0.001
	(0.001)*	(0.251)**	(0.000)***
Dutch guilder	0.031	1.003	0.002
	(0.032)***	(0.321)*,(1)	(0.000)***
Finnish markka	0.025	2.883	0.007
	(0.008)***	(0.777)***	(0.000)***
French franc	0.050	1.777	0.002
	(0.014)***	(0.467)***	(0.000)***
Greek drachma	0.023	1.238	0.007
	(0.007)***	(0.290)***	(0.000)***
Irish punt	0.010	1.121	0.050
	(0.007)	(1.082)	(0.003)***
Italian lira	0.126	0.537	0.005
	(0.045)***	(0.335)	(0.000)***
Portuguese escudo	0.024	0.395	0.003
	(0.010)**	(0.277)	(0.000)***

Notes : Standard errors are in parantheses. *, **, *** denote statistical significance at 10%, 5% and 1% levels. (1) The bootstrapped standard error is reported.